The Health Plan Choices of Retirees Under Managed Competition

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Objective. To investigate the effect of price on the health insurance decisions of Medicare-eligible retirees in a managed competition setting.

Data Source. The study is based on four years of administrative data from the University of California (UC) Retiree Health Benefits Program, which closely resembles the managed competition model upon which several leading Medicare reform proposals are based.

Study Design. A change in UC's premium contribution policy between 1993 and 1994 created a unique natural experiment for investigating the effect of price on retirees' health insurance decisions. This study consists of two related analyses. First, I estimate the effect of changes in out-of-pocket premiums between 1993 and 1994 on the decision to switch plans during open enrollment. Second, using data from 1993 to 1996, I examine the extent to which rising premiums for fee-for-service Medigap coverage increased HMO enrollment among Medicare-eligible UC retirees.

Principle Findings. Price is a significant factor affecting the health plan decisions of Medicare-eligible UC retirees. However, these retirees are substantially less price sensitive than active UC employees and the non-elderly in other similar programs. This result is likely attributable to higher nonpecuniary switching costs facing older individuals.

Conclusions. Although it is not clear exactly how price sensitive enrollees must be in order to generate price competition among health plans, the behavioral differences between retirees and active employees suggest that caution should be taken in extrapolating from research on the non-elderly to the Medicare program.

Key Words. Managed competition, price elasticity, health plan choice, switching costs.

A number of leading Medicare reform proposals would restructure the program based on the principles of "managed competition" (Enthoven 1988; Aaron and Reischauer 1995; Butler and Moffit 1995; Kendall 1995; Dowd, Feldman, and Christianson 1996; Cutler 1997). Advocates of this approach often point to large employer-sponsored health benefits programs as examples of how such a program would work in practice. For example, Butler

and Moffit (1995) hold up the Federal Employees Health Benefit Program as a model, and Fitzgerald (1995) argues that lessons can be learned from the experiences of large corporations. Dowd, Feldman, and Christianson (1996) refer extensively to the health benefits programs for public employees in Minnesota and Wisconsin and to their own research on the plan choices of employed individuals (Feldman et al. 1989; Feldman and Dowd 1993; Dowd and Feldman 1994/1995). These and other recent studies (Long, Settle, and Wrightson 1988; Buchmueller and Feldstein 1997; Royalty and Solomon 1999; Cutler and Reber 1998) indicate that employees are willing to switch plans in response to small changes in out-of-pocket premiums. Where a fixed dollar contribution policy has been implemented, there is evidence that it has led to lower premiums (Feldman and Dowd 1993; U. S. GAO 1994; Enthoven and Singer 1996; Cutler and Reber 1998; Buchmueller 1997).

Such results make managed competition an attractive model for Medicare reform. However, the extent to which the experience of large employer-sponsored benefit programs generalizes to Medicare is unclear. Because of their greater use of medical care, Medicare beneficiaries are likely to face substantially higher "switching costs" than non-elderly workers.[1] In addition, research on how the elderly perceive various health insurance options suggests they place much more importance on other factors, such as quality of care, freedom of referral, and the burden of paperwork, than on premiums (Harris 1997).

In this study, I examine the health plan choices made by Medicare beneficiaries in a setting that closely resembles leading proposals for Medicare reform. While there are numerous studies of the health insurance decisions of Medicare beneficiaries, this is the first to look at choices made under the conditions of managed competition and also the first to estimate directly the effect of out-of-pocket premiums.[2] The data come from the University of California's (UC) health benefit program. UC retirees face a choice of plans that includes a fee-for-service (FFS) plan supplementing basic Medicare (i.e., a "Medigap" plan) along with several HMOs. Recent changes in the UC

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program provide a natural experiment for investigating how Medicare beneficiaries would respond to price changes brought about by market-oriented reforms. Prior to 1994, none of the available plans required a premium contribution from Medicare-eligible retirees. Because of a reduction in UC's contribution, the FFS plan and several of the HMOs began in 1994 to require monthly payments from enrollees. Out-of-pocket premiums for FFS coverage increased again in 1995 and 1996.

I use data on the open enrollment choices made by Medicare-eligible retirees from 1993 to 1996 to investigate two related research questions. First, I examine how the price changes that occurred between 1993 and 1994 as a result of the policy change affected the decision to switch plans. Because benefits and other plan features remained constant, the relationship between changes in out-of-pocket premiums and plan switching provides a clean estimate of the price sensitivity of Medicare-eligible retirees. Next I use data from the full four-year period to estimate the effect of price on the demand for FFS coverage. This second analysis provides direct evidence on the extent to which Medicare managed care enrollment would increase if beneficiaries were exposed to the difference in cost between FFS and HMO coverage.

The two analyses generate similar qualitative results. Both indicate that price is a significant, although small, factor influencing the health plan choices of Medicare-eligible retirees. Comparisons with previous studies suggest that retirees are much less sensitive to price than non-elderly individuals choosing plans under similar circumstances.

MANAGED CARE AND HEALTH PLAN CHOICE IN TODAY'S MEDICARE

When the Medicare program was established in 1965, its benefit design resembled the private insurance policies that were typical at the time. Since then, the private insurance market has been completely transformed. Indemnity insurance–paying providers on an FFS basis have been supplanted by HMOs, preferred provider organizations (PPOs), and a variety of hybrid managed care plans. While there have been significant changes in the way Medicare pays providers—e.g., the introduction of the Prospective Payment System for hospitals in 1983 and the Resource-Based Relative Value Scale for physicians in 1992—from the beneficiaries' perspective, basic Medicare has changed very little over the past 30 years. An important qualification to this statement is that since the mid-1980s, beneficiaries have had the option

of enrolling in an HMO instead of basic Medicare. While Medicare HMO enrollment has grown steadily over time, the percentage of beneficiaries in HMOs remains relatively low and highly concentrated in a small number of markets (Zarabozo, Taylor, and Hicks 1996; Welch 1996).

HMOs participating in the Medicare program must provide a basic set of benefits, which includes those covered under FFS Medicare, and may also offer supplemental benefits. While HMOs may be able to provide care for less than the amount the government pays them, they are prohibited from rebating savings to beneficiaries. As a result, rather than lowering prices, plans compete by offering additional benefits (Feldman et al. 1993). Prescription drug coverage and other benefits not covered under basic Medicare are particularly attractive to many beneficiaries. Medicare HMO enrollees have substantially lower out-of-pocket medical expenses than individuals with only basic Medicare coverage and lower premium payments than those with basic Medicare and a private supplemental policy.[3]

The introduction of HMOs to Medicare was intended to reduce program spending, although it appears to have had the opposite effect. The reason has to do with the way payments to HMOs are based on costs in the FFS sector. For every beneficiary enrolling in an HMO, Medicare pays 95 percent of the average cost for similar individuals covered by FFS Medicare in the same area. While there is some attempt to adjust for differences in risk, it is widely believed that Medicare HMO enrollees are less costly to insure than FFS beneficiaries. Thus, the 95 percent payment overstates the cost of providing care to HMO enrollees and thereby increases total Medicare spending (Brown, Clement, Hill, et al. 1993).

COMPETITION-BASED PROPOSALS FOR MEDICARE

There is general agreement among health policy analysts on the need to break the link between FFS costs and what Medicare pays HMOs, and to modify other distortionary aspects of the current system.[4] Several recent Medicare reform proposals would do this as part of reorganizing the program to encourage cost-conscious choices by beneficiaries and price competition among health plans and providers.

Dowd, Feldman, and, Christianson (1996) lay out a detailed program that is fairly representative of these market-oriented proposals. In their "competitive pricing" model, Medicare recipients would choose from a menu consisting of all health plans in an area that meet minimum qualification standards

and that agree to offer a specified benefits package. Plans would be allowed to offer supplemental benefits in addition to the standard package. Beneficiaries would choose plans during an annual open enrollment period, and the ability to switch plans at other times would be restricted. The government would contribute a fixed amount based not on costs in the FFS sector, as under the current system, but on the premium of the area's lowest-cost plan. Individuals choosing a more expensive plan would pay the full difference in premiums.

Other market-oriented proposals differ slightly from Dowd, Feldman, and Christianson's in terms of such things as the extent to which benefits should be standardized, how exactly the government's contribution should be set, and the need for risk-adjusting payments to plans. Overall, however, these differences are less significant than the similarities. The basic premise of all these proposals is that changing the incentives facing beneficiaries and plans will enhance economic efficiency and reduce the growth in Medicare spending. An additional motivation for reforming Medicare in this way is that the program is an anachronism in today's health care system, dominated as it is by managed care. That is, apart from arguments related to economic efficiency, an additional argument for these proposals is that they would make Medicare benefits comparable to those received by non-elderly Americans with private insurance.

THE UNIVERSITY OF CALIFORNIA RETIREE HEALTH BENEFITS PROGRAM

The health benefits offered to the employees and retirees of the University of California closely resemble the reform models proposed by Dowd, Feldman, and Christianson (1996) and others. The UC consists of nine campuses located in California and three national laboratories, including the Los Alamos National Laboratory in New Mexico. In 1996, roughly 90,000 active employees and 33,000 retirees were eligible for health benefits. In this study, I focus on the behavior of roughly 17,000 retirees who are either themselves eligible for Medicare or whose spouse is Medicare eligible.

Just like the various Medicare reform proposals, there is an annual open enrollment period during which UC retirees choose from a menu consisting of several HMOs and an indemnity plan that supplements FFS Medicare coverage. As is typical of employer-sponsored programs, UC retirees are offered the same plan choices as active employees (Morrisey, Jensen, and Henderlite 1990). As would be required under some, but not all reform

proposals, the HMOs provide an essentially standard set of benefits.[5] The FFS plan features a "coordination of benefits" design, whereby it pays Medicare deductibles and coinsurance, leaving retirees with essentially no out-of-pocket costs for Medicare covered services.[6] In addition, the plan pays for things not covered by Medicare, most notably prescription drugs.[7]

Another important similarity with the managed competition proposals is that the UC makes a fixed-dollar contribution toward each enrollee's premium. Through 1993, the contribution was based on a weighted average of the active employee premiums charged by the four plans with the highest UC enrollment. Because this group included a "high option" indemnity plan, the average was higher than the premium of all the HMOs. Financial pressure brought on by the California state budget crisis of the early 1990s led to a change in the UC's contribution policy. In January 1994, the UC reduced its contribution to equal the active employee premium for the least expensive plan available statewide. Employees and retirees must now pay the full difference between this amount and the premium for their chosen plan.

Because premiums for retirees represent either the cost of coverage that supplements basic Medicare (in the case of the FFS plans) or the incremental cost of HMO benefits beyond the standard Medicare plan, they are substantially lower than the premiums charged for active employees. Despite this, the UC contributes the same amount for retirees as it does for active employees. As a result, even after the policy change, the UC's contribution exceeded the gross retiree premium for most plans. However, the change did cause monthly out-of-pocket costs to increase for retirees in certain plans. The reason is that the UC allows retirees to apply the difference between the UC contribution and the premium for their chosen plan to their Medicare Part B premium.[8] Beginning in 1994, retirees in certain plans were required, for the first time, to pay a portion of their Part B premium.[9]

Table 1 summarizes the health plan choices of UC retirees in 1993, the year before the change in the contribution policy. Separate tabulations are presented for retirees in northern California, southern California, and New Mexico to account for differences in plan offerings and market conditions across the regions. Since California has the highest rate of managed care penetration in the country among both the non-elderly and Medicare beneficiaries, it is not surprising that in 1993 a large fraction of UC retirees were already in HMOs. In 1994, Riverside County, CA had the highest Medicare HMO market share in the country, with 47 percent of beneficiaries in HMOs. Three other southern California counties, San Diego, Orange, and Los Angeles, ranked second, fourth, and eighth with HMO market shares of

Table 1: The University of California Retiree Health Benefits Program: 1993 Enrollment by Region and Health Plan

Plan Name	*Plan Type*	*10/93 Enrollment* N (%)	*1993 OOP Premium* Sing./2-Party	*1994 OOP Premium* Sing./2-Party	*% of 1993 Enrollees Switching in 1994*
A. Northern California					
Pru High Option	FFS Medigap	4452 (45.8)	0/0	33/62	1.7%
Kaiser North	Group HMO	3119 (32.1)	0/0	0/0	1.1%
Qual Med (CA)	Network HMO	634 (6.5)	0/0	17/30	6.4%
Foundation	Network HMO	622 (6.4)	0/8	35/65	5.4%
Health Net	Network HMO	150 (1.5)	0/0	0/0	0.7%
Take Care	Network HMO	114 (1.2)	0/0	22/40	7.0%
B. Southern California					
Pru High Option	FFS Medigap	4535 (66.7)	0/0	33/62	1.8%
Kaiser South	Group HMO	1206 (17.7)	0/0	0/0	2.0%
Health Net	Network HMO	853 (12.7)	0/0	0/0	2.0%
UC Care	PPO	112 (1.7)	0/0	0/0	20.2%
C. New Mexico					
Los Alamos Plan	FFS Medigap	1648 (89.6)	0/5	65/126	24.1%
Lovelace	HMO	138 (7.5)	0/0	0/0	0.0%
Qual Med (NM)	HMO	31 (1.7)	0/0	18/32	8.5%

Note: Monthly out-of-pocket premium = max [0, gross plan premium + Medicare Part B premium − UC contribution]. Two-party premiums reported are for couples where both husband and wife are Medicare-eligible. Northern California locations are Berkeley, Davis, Livermore, Santa Cruz, and San Francisco. Southern California locations are Irvine, Los Angeles, Riverside, San Diego, and Santa Barbara. Market share percentages for northern and southern California do not add to 100 because plans with very small enrollments (including UC Care in northern California) and those terminated between 1993 and 1994 are not listed.

41 percent, 38 percent, and 33 percent, respectively (Welch 1996). The UC has campuses in all four of these counties and their ranking in terms of HMO enrollment parallels the general market: HMO market share is highest at UC Riverside, followed by UC San Diego, UC Irvine (in Orange County), and UCLA. Among all southern California campuses (these four plus UC Santa Barbara), 32 percent of retirees were in HMOs in 1993. The percentage of retirees enrolled in HMOs was even higher, roughly 50 percent, in northern California. While two HMOs were available at Los Alamos, NM, in 1993 they enrolled only 8 percent of the retirees at this location.

THE EFFECT OF PRICE ON HEALTH PLAN SWITCHING BY RETIREES

The third and fourth columns of Table 1 show how the reduction in the UC's contribution affected monthly out-of-pocket premiums for all plans.[10] Premiums increased the most for the two indemnity plans, and also rose for three HMOs in northern California and one of the two HMOs offered in New Mexico. Five HMOs and the one PPO plan remained free to UC retirees in 1994. Since benefits remained constant and there were no other significant changes to the health plan options of UC retirees, the reduction of the UC's premium contribution creates a good natural experiment for investigating the price sensitivity of Medicare recipients in a managed competition setting. We can infer the effect of price on plan switching by comparing UC retirees who faced an increase in out-of-pocket premiums between 1993 and 1994 with a control group of retirees who faced constant (zero) out-of-pocket premiums.

The percent of each plan's 1993 enrollees who voluntarily switched to another plan between 1993 and 1994 is reported in the last column. The response of FFS enrollees to the increase in premiums differs substantially between the California and New Mexico locations. At Los Alamos, where initial HMO penetration was low and where FFS premiums increased to between $65 and $155, 24 percent of FFS enrollees switched plans. In contrast, while the monthly employee premium for Prudential High Option increased from zero to $33 for single ($62 for two-party) coverage, less than 2 percent of retirees in the plan switched out.

A comparison of switching rates for the HMOs that did and did not experience premium increases indicates a small but statistically significant response to price. Out-of-pocket premiums for three plans, Qual Med, Foundation, and Take Care increased by an average of $34 per month. Hereafter,

these plans will be referred to as "pay" HMOs, in contrast to the "free" HMOs whose premiums remained constant at zero. The switching rate for the pay HMOs as a group was 6 percent. Switching rates were lower for the free HMOs. The most appropriate comparison plan is Health Net, which has a similar design as the three pay plans and, in many locations, contracts with the same providers. Whether Health Net enrollees in northern California or those statewide are used as the comparison group, the difference in their switching rate and the rate for the pay HMOs is statistically significant at the one- percent level. When the two Kaiser plans are used as a comparison group, the difference in switching rates is even more pronounced. However, Kaiser represents a less suitable control group due to differences between network and group model HMOs: switching from Kaiser requires switching physicians, whereas switching among the network plans often does not.

AN ECONOMETRIC MODEL OF PLAN SWITCHING

Because gross premiums differ across plans as well as across coverage categories, there was a fair amount of variation in price increases between 1993 and 1994. To account fully for this variation and to control for other factors, I estimate the following probit regression model of the switching decision:

$$S_i^* = X_i\beta + \gamma\Delta P_i + \varepsilon_I \tag{1}$$

$$S_i = 1 \text{ if } S_i^* > 0$$

$$= 0 \text{ otherwise,}$$

where S^* is the latent propensity to switch plans, and the indicator variable S equals one for individuals who switched plans during the 1993 open enrollment period (when 1994 plans were chosen). X is a vector of controls and ΔP_i is the change between 1993 and 1994 in out-of-pocket premiums for the plan chosen by individual i in 1993.

Similar to the univariate tests described above, this regression compares the switching behavior of retirees facing price increases of various magnitudes to those in plans that remained free in 1994. Accordingly, the absolute change in premiums can be thought of as a change in relative prices under the assumption that retirees compared the utility of remaining in their original plan to the utility of switching to a free plan. The fact that over 80 percent of plan switchers moved into one of those free plans provides some support for

this assumption. In other studies where there is not such an obvious control group, failure to account for changes in other plans' prices may be a source of bias.[11]

There are several reasons why I focus on the switching decision, rather than estimate a model of health plan choice. First, standard choice models assume an equivalence between an individual's initial choice and the subsequent decision to switch plans during open enrollment. However, if switching costs are large, even individuals who were fairly price sensitive initially may be less sensitive to price changes later. The potential for such persistence suggests the value of examining directly the decision to switch plans. Second, premiums may be correlated with plan and enrollee characteristics, which are important determinants of the choice decision but which are not observed by the analyst. Unobservables that are relatively constant over time will be a source of bias for models defined in price *levels*, but less of a problem for models based on price *changes*, as their effect is essentially differenced away.

That is not to say that the problem of omitted variable bias can be ruled out completely, as unobserved factors may also be related to the propensity to switch plans. The most important shortcoming of the UC data set is that it includes no information on enrollees' health status. Omitting health measures from the switching regression will bias the coefficient on ΔP if (a) health status has a direct effect on switching and (b) plans that increased in price between 1993 and 1994 were ones that had previously been particularly attractive or unattractive to individuals in poor health.[12] Plausible theoretical arguments can be made for either a negative or positive effect of poor health on switching. On one hand, sicker enrollees are more likely than their healthier counterparts to have strong ties to their personal physicians and therefore may be less likely to change plans if doing so requires changing providers. By the same logic, however, switching may be positively correlated with poor health in settings where plans' provider panels change significantly from year to year. In addition, as greater users of the system, less healthy individuals are likely to be more demanding as consumers and perhaps more likely to switch plans in response to unsatisfying experiences.

The previous literature provides limited insight into the direct effect of health status on the decision to switch plans. While some early studies find a negative relationship between poor health (or prior utilization) and disenrollment (Wersenger and Sorenson 1982; Hennelly and Boxerman 1983; Griffith 1984), several others, including the two studies most closely related to this one (Long, Settle, and Wrightson 1988; Riley, Feuer, and Lubitz 1996), find little relationship.[13]

In the UC data, there is likely to be a negative correlation between health status and ΔP in the full sample because premiums increased the most for the FFS plans, and prior research suggests that Medicare beneficiaries with FFS coverage are less healthy than those in HMOs (Brown, Clement, Hill, et al. 1993). However, health status is not likely to be correlated with ΔP among HMO enrollees. The HMOs available to UC retirees are required to have the same benefits, and among the network model plans there is a high degree of provider-panel overlap. Because of these similarities and the fact that all the HMOs were free to retirees prior to 1994, there is little reason to expect that individuals in poor health were either more or less likely to be in a plan that later went up in price. Therefore, when the sample is restricted to HMO enrollees, the coefficient on ΔP should not be biased due to the omission of health-related controls.

SWITCHING MODEL RESULTS

Switching model results are presented in Table 2. The results in column 1 pertain to all UC retirees who were either covered by Medicare or whose spouse was covered, and whose 1993 plan was available in 1994.[14] Results in the second column are for retirees who were initially in an HMO. Both samples exclude retirees who were living out-of-state in 1993. The rationale for this exclusion is that only a subset of the plans offered by the UC will be viable options for these individuals.[15] Summary statistics for all variables are in the appendix, Table A1.

The results indicate that the probability of switching plans decreases with age, although in the HMO sample the effect is not statistically significant. This is consistent with the notion that switching costs increase with medical utilization and agrees with the results of several previous studies on plan switching by the non-elderly.[16] For both the full sample and the HMO-only sample, there are no significant differences in switching rates by coverage category or, among single individuals, by gender. Most of the location coefficients are insignificant, although as a group they are significant for each sample.

The results indicate a statistically significant positive relationship between increases in out-of-pocket premiums and plan switching, though the magnitude of the effect is fairly small. The probability derivative for ΔP is reported in brackets, although the price effect is best illustrated by using the estimated coefficients to simulate the probability of switching plans in

Table 2: The Effect of Price on Health Plan Switching: Probit Regression Results

Independent Variables	*(1) All Plans*	*(3) HMOs Only*
Change in monthly premium (ΔP)	.0073***	.0200***
	(.0027)	(.0030)
	[.0004]	[.0008]
Age	−.029***	−.015
	(.006)	(.010)
Male with single coverage	.208	.181
	(.167)	(.113)
Female with single coverage	.085	.053
	(.160)	(.134)
Couple with non-Medicare member	.145	.169
	(.135)	(.144)
HMO in 1993	.130	
	(.192)	
S.F. Bay Area	−.024	−.277***
	(.085)	(.119)
UC Davis	−.071	−.347***
	(.130)	(.126)
UC Irvine	−.076	.021
	(.107)	(.203)
UC Riverside	−.015	−.390*
	(.138)	(.205)
UC San Diego	−.046	−.085
	(.103)	(.158)
UC Santa Cruz	.213**	.101
	(.096)	(.289)
UC Santa Barbara	−.059	.033
	(.193)	(.289)
Lawrence Livermore National Lab.	−.134*	−.558***
	(.095)	(.203)
Los Alamos National Lab.	.941***	−.309
	(.157)	(.286)
Constant	−.285	−.997
	(.544)	(.771)
Number of observations	16,104	6,928
Log-likelihood	−2,298	−679

Note: Dependent variable equals one if individual switched plans between 1993 and 1994, zero otherwise. Robust standard errors in parentheses. Probability derivatives (evaluated at the mean of each sample) are in brackets. Bay Area locations are UC Berkeley, Lawrence Berkeley Laboratory, UC San Francisco, and Hastings School of Law. UCLA is the omitted location.

*Statistically significant at the .10 level; **statistically significant at the .05 level; ***statistically significant at the .01 level.

response to various price increases. For each individual in the sample, I predict the probability of switching plans when premiums are constant and when they increase by between $10 and $60 per month, and then average the predicted values over the sample. These simulation results are reported in Table 3.

The full sample model predicts that 2.3 percent of enrollees will switch plans when premiums are constant. An increase of $20 per month will raise the switching rate slightly to 3.0 percent. Premium increases of $40 and $50 result in switching rates of 3.9 percent and 4.4 percent, respectively. As would be expected based on the figures in Table 1, the results for the HMO-only sample indicate a stronger price response. The simulations imply that a $20 price increase will raise the switching rate for HMO enrollees to 3.4 percent from a baseline of 1.4 percent. The HMO-only switching rate is 10.8 percent when $\Delta P = \$50$.

In this particular setting, FFS enrollees may be less price sensitive than HMO enrollees because they are less likely to view the available free plans as good substitutes for their original plan. In addition, the difference between the full sample results and those for the HMO-only sample may be related to the fact that the model does not control for health status. As noted, if the propensity to switch plans is negatively correlated with poor health, and retirees with FFS coverage in 1993 are on average less healthy than those

Table 3: The Effect of Price on Switching Among Health Plans: Simulated Switching Probabilities

	(1) *All Plans*	*(3)* *HMOs Only*
Probability of Switching Plans if . . .		
$\Delta P = \$0$	.023	.014
$\Delta P = \$10$	.026	.022
$\Delta P = \$20$	.030	.034
$\Delta P = \$30$	.034	.052
$\Delta P = \$40$	.039	.076
$\Delta P = \$50$	.044	.108
$\Delta P = \$60$	.050	.149

Note: Simulated probabilities are based on the parameters reported in Table 2. For each observation, ΔP is set to 0, 10, 20, 30, 40, 50, and 60, and the probability of switching is predicted at each level. Predicted probabilities are then averaged over the estimation sample.

in HMOs, the price response for the full sample will be biased toward zero. This is much less of a concern for the HMO-only sample.

One way to put these results in perspective is to compare them to results for active UC employees. This comparison is salient given that predictions concerning the effect of transforming Medicare into a managed competition program are often based on evidence from non-elderly individuals in large employer-sponsored groups. Active UC employees are a good comparison group because they choose from the same menu of plans as UC retirees, and faced similar changes in out-of-pocket premiums between 1993 and 1994.

Regression results for the full sample of active employees predict that when premiums are constant, 5 percent of active UC employees will switch plans. When out-of-pocket premiums increase by $20, models estimated using active employee data predict switching rates of 30 percent when all plans are considered, and 34 percent when only HMO enrollees are analyzed.[17] Thus, compared to active UC employees, retirees were much less likely to switch plans in general and were much less responsive to changes in out-of-pocket premiums.

One potential explanation for this result is that sensitivity to price may decrease with enrollees' age. To investigate this possibility, I stratified the active employee and retiree samples by age and estimated separate switching regressions on the stratified samples. Consistent with the results in Table 2, these regressions, which are not reported, indicate a strong negative relationship between age and the probability of switching plans, holding price constant. The comparison is clearest when attention is restricted to individuals who were initially in HMOs. For this group, the simulated switching rate when $\Delta P = 0$ for older active employees (over the age of 55) is 3.2 percent, which lies roughly halfway between a baseline rate of 5.9 percent for employees under age 35 and a baseline rate of 0.8 percent for younger UC retirees under age 73. However, there is a quite different pattern across these groups in terms of estimated price effects. Among active employees, the probability derivative with respect to price is larger for those under age 35 than for those 55 and older, although the difference is small (.0073 versus .0066) and not statistically significant. In contrast, the probability derivative for retirees under age 73 is substantially lower (.0009). The price effect is even smaller for retirees age 73 and older (.0004). Thus, while price sensitivity does decline with age, age does not appear to be the primary factor explaining the large difference between active employees and Medicare-eligible retirees.[18]

THE EFFECT OF OUT-OF-POCKET PREMIUMS ON THE DEMAND FOR FFS MEDICARE

While it is not an explicit goal of managed competition reform proposals to increase the number of Medicare beneficiaries enrolled in managed care plans, many analysts believe such an outcome is likely. This belief is based on the twin assumptions that managed care plans will be more effective than FFS Medicare in controlling cost growth, and that when these cost differentials are reflected in premiums, many beneficiaries currently enrolled in traditional FFS Medicare will switch to less costly HMOs. The switching results from the 1993 Open Enrollment period provide indirect evidence on the extent to which exposing Medicare beneficiaries to the higher cost of FFS coverage will cause HMO enrollment to increase. In this section, I use data from the period 1993 through 1996 to address the question more directly.

Table 4 provides summary information on the premiums faced by UC retirees for FFS and managed care plans, and the FFS market share for the years 1993 through 1996. As shown in Table 1, between 1993 and 1994 prices increased the most for FFS coverage at the New Mexico location. In 1995, choice was eliminated and all employees and retirees were put into a single point-of-service (POS) plan. FFS Medigap coverage remained an option for UC retirees in California for the entire period. There, Prudential premiums increased by between 22 percent and 28 percent, depending on coverage tier, between 1994 and 1995, and by comparable percentages between 1995 and 1996. In all three regions, the increasing FFS/HMO price differential coincided with a decline in FFS market share.

To estimate more precisely the effect of out-of-pocket premiums on the demand for FFS coverage, and to control for other factors, I estimate several regressions. Feldman et al. (1989) propose a nested logit model for this type of estimation problem. In their analysis using data from the early 1980s, they divide health plans into two nests, one for group and staff HMOs, the other for IPA-model HMOs and FFS plans. In the UC data, a natural classification scheme would combine the HMOs and the one PPO/POS plan into one nest, and to consider the one FFS plan at each campus to be a second nest.

One problem with estimating such a model is that not only is there just one plan in the FFS nest, but there is little variation within the managed care nest. As noted, the benefits provided by all the HMOs are virtually identical by design. The variation in price among the managed care plans is also fairly limited. None of the HMOs offered in southern California required retirees to

Table 4: Premiums and Market Share by Region and Plan Type, 1993–1996

	1993	1994	1995	1996
A. Northern California				
Number of FFS plans	2[a]	1	1	1
Mean premium ($ per month)	2.83	50.05	61.76	81.60
Number of managed care plans[b]	7	7	5	6
Mean premium ($ per month)	0.45	9.99	5.78	4.80
FFS market share	50.5%	44.5%	42.9%	40.9%
B. Southern California				
Number of FFS plans	1	1	1	1
Mean premium ($ per month)	3.71	51.39	63.33	85.58
Number of managed care plans[b]	4	4	4	5
Mean premium ($ per month)	0.08	0.22	3.04	3.87
FFS market share	66.6%	63.0%	61.9%	59.4%
C. New Mexico				
Number of FFS plans	1	1		
Mean premium ($ per month)	2.34	101.87		
Number of managed care plans	2	2		
Mean premium ($ per month)	0	4.23		
FFS market share	90.6%	69.1%		

Note: Mean premiums represent the average amounts actually paid by enrollees; therefore, differences across locations represent differences in coverage tier. Choice was eliminated in New Mexico after 1994, thus data on premiums and enrollments from that site are not relevant in 1995 and 1996.

[a] Two FFS plans were offered at UC Davis; neither required premium contributions.

[b] In 1993 and 1994 the managed care options consisted of HMOs plus one PPO. In 1995, the PPO was changed to a point-of-service plan.

make out-of-pocket premium contributions over this period, and premiums for the pay HMOs in northern California changed very little between 1994 and 1996. In contrast to the lack of variation within nests, there is substantial price variation across nests. As shown in Table 4, the price of FFS coverage varies significantly over time. Differences between FFS plans in California and New Mexico provide additional cross-sectional variation in price, as does the fact that out-of-pocket premium contributions differ according to coverage type and the Medicare eligibility of family members. Thus, while it may not be possible to identify the effect of price in a multinomial model

that incorporates the full set of options available to each retiree, it is possible to estimate a binary model of whether a retiree chooses FFS coverage rather than one of the managed care plans (i.e., the choice between the two nests). Such a model directly addresses the policy question of whether rising FFS costs will increase managed care enrollment.

The model I estimate is

$$F_{it}^{*} = X_{it}\delta + \theta P_{it}^{F} + u_{it} \tag{2}$$

$$F_{it} = 1 \text{ if } F_{it}^{*} > 0$$

$$= 0 \text{ otherwise,}$$

where the latent variable F_{it}^{*} represents the demand for FFS coverage; its observable analog, F_{it}, equals 1 if in year t individual i chose the FFS plan available at her location, and zero if she chose a managed care plan. P^{F} is the out-of-pocket premium for the FFS plan and X includes the same demographic control variables used in the switching analysis.

The error term in Equation 2 can be decomposed as

$$u_{it} = \zeta_i + \nu_{it}, \tag{3}$$

where ζ_i represents individual i's time-invariant taste for FFS coverage and ν_{it} is a normally distributed mean zero error term. Given the sources of price variation in my data, there is no reason to suspect ζ to be correlated with P^{F}.[19] Therefore, I estimate Equation 2 using a conventional probit model. Then, as in the switching analysis, I adjust the standard errors to account for clustering in the data. This causes the estimated standard errors to increase considerably relative to the case where the errors are assumed to be i.i.d.

The regression coefficients and standard errors are presented in Table 5. For the price variable I also report (in brackets) the probability derivative evaluated at the sample mean. In column 1 the model is estimated for all UC retirees; in columns 2 through 4, the sample is stratified by region (northern California, southern California, and New Mexico). As in the switching analysis, I exclude retirees living out of state based on the rationale that HMO coverage may not be a viable option for retirees who have moved away from their former place of employment.[20] Summary statistics for the full sample are reported in the appendix.

As expected, the demand for FFS coverage increases with age. When the regression function is evaluated at the sample means, the age coefficient in column 1 implies that the probability of FFS coverage increases by 2

Table 5: The Effect of Price on the Demand for Fee-for-Service Medicare: Probit Regression Results

Independent Variables	*(1) All Locations*	*(2) Northern California*	*(3) Southern California*	*(4) New Mexico*
Premium for FFS plan (P^F)	−.0033***	−.0033***	−.0024***	−.0083***
	(.0004)	(.0004)	(.0003)	(.0007)
	[−.0013]	[−.0013]	[−.0010]	[−.0025]
Age	.050***	.049***	.055***	.030***
	(.001)	(.001)	(.002)	(.007)
Male with single coverage	−.151***	−.084***	−.261***	−.174
	(.037)	(.038)	(.028)	(.163)
Female with single coverage	−.077**	−.035	−.135***	.032
	(.032)	(.037)	(.030)	(.128)
Couple with non-Medicare member	−.047**	−.029	−.126***	−.186**
	(.038)	(.046)	(.034)	(.100)
S.F. Bay Area	−.663***	.103		
	(.027)	(.065)		
UC Davis	−.765***			
	(.058)			
UC Irvine	.107***		.102***	
	(.031)		(.030)	
UC Riverside	−.869***		−.877***	
	(.019)		(.019)	
UC San Diego	−.235***		−.234***	
	(.010)		(.010)	
UC Santa Barbara	−.488***		−.498***	
	(.032)		(.033)	
UC Santa Cruz	−.163	.599***		
	(.034)	(.037)		
Lawrence Livermore Lab.	−.873***	−.103		
	(.036)	(.076)		
Los Alamos National Lab.	.253*			
	(.141)			
Constant	−2.999	−3.661	−3.296	−.937
	(.096)	(.143)	(.108)	(.473)
Number of observations	68,888	39,214	26,705	2,969
Log-likelihood	−42,333	−24,623	−16,099	−1,494

Note: Data for the full sample and California subsamples are for the period 1993 to 1996; data from New Mexico is for 1993 and 1994. The dependent variable equals 1 for retirees choosing FFS Medigap coverage, 0 otherwise. Robust standard errors are in parentheses. Probability derivatives (evaluated at the mean of each sample) are in brackets. The omitted coverage category is married couples in which both the husband and the wife are Medicare-eligible. Bay Area locations are UC Berkeley, Lawrence Berkeley Laboratory, UC San Francisco, and Hastings School of Law. In columns 1 and 3, UC Irvine is the omitted location; in column 2, UC Davis is the omitted location.

*Statistically significant at the .10 level; **statistically significant at the .05 level; ***statistically significant at the .01 level.

percentage points for every year. In New Mexico, all the coverage category coefficients are insignificant. In the full sample and the two California regions, single men are least likely to choose FFS coverage, and married couples where both husband and wife are Medicare-eligible (the omitted category) are most likely to opt for FFS coverage.

For all samples, the coefficient on the FFS out-of-pocket premium is negative and statistically significant at the one- percent level. Evaluated at sample means, the full sample results imply that a $10 increase in the monthly cost of FFS Medigap coverage will reduce the percentage of retirees choosing such coverage by 1.3 percent, which corresponds to a demand elasticity of −.16. The elasticity is slightly larger for the northern California sample (−.19) and slightly smaller for the southern California sample (−.09). This difference is due to a slightly stronger price effect and a lower average FFS market share in northern California. The estimated elasticity is highest for the New Mexico sample, where the marginal effect of price is −.0025. Evaluated at the New Mexico sample means, this implies an FFS price elasticity of −.22; combining the New Mexico price coefficient with the northern California sample means for P^F and F produces an estimated elasticity of −.35.

The results are robust to changes in sample definition. When I exclude couples in which one spouse is not eligible for Medicare, the probit results imply that a $10 increase in the monthly FFS premium will reduce the FFS market share by 1.4 percent. Models estimated separately for single individuals and those covering dependents indicate that single individuals are somewhat more price sensitive, although the difference is small, and for both groups the negative effect of price is statistically significant at the one- percent level. I also estimated models that allowed the effect of price to vary across three age categories—under 70, ages 70 to 79, and ages 80 and up. As might be expected, the results indicate that price sensitivity decreases with age. The partial derivative with respect to price for the youngest age category is nearly twice the magnitude of that for the oldest category,−.00133 versus −.00071.

As with the switching analysis, these results can be put in perspective by comparing them to elasticities estimated for non-elderly employed individuals. Three recent studies using data from similar employer-sponsored settings and similar time periods provide useful comparisons. All three generate price elasticities larger than those just reported.

Dowd and Feldman (1994/1995) use aggregate data to estimate premium elasticities for employees from five Twin Cities firms. At the time of their analysis (1989–1993), the Twin Cities, like California, was a mature managed care market. At the means for their sample, Dowd and Feldman's

results imply an "insurer perspective" elasticity of −7.9. When calculated using the out-of-pocket price faced by enrollees, the estimated elasticity is roughly −1. Evaluated at market shares relevant to the FFS plans in my sample, the "enrollee perspective" elasticity falls further, to between −0.5 and −0.7. Royalty and Solomon (1999) analyzed the health plan choices of Stanford University employees in 1994 and 1995. Not only does the time period coincide with this study, but there is some overlap in the plans offered to Stanford employees and UC retirees in northern California. Royalty and Solomon's multinomial logit results imply price elasticities ranging from −0.4 to −0.8. Parameter estimates from a fixed effect logit model imply even larger elasticities. Finally, Cutler and Reber (1998) also examine the health plan choices of university employees, in this case Harvard. The estimation problem they face is similar to the one in this study. While the Harvard menu included several plans, in the years Cutler and Reber analyzed it (1994 to 1996), the variation in price among HMO offerings was quite limited. Thus, rather than estimate a full choice model, they estimate a binary logit model of whether employees chose the single PPO option or one of several HMO options. Their estimates imply out-of-pocket premium elasticities of between −0.3 and −0.6.

DISCUSSION

In this study, I present two separate but related models of decisions made by Medicare beneficiaries in a setting that resembles leading Medicare reform proposals. The first focuses on the willingness of retirees to switch health plans when the price of their current plan rises relative to similar alternatives. The importance of this question derives from the fact that the managed competition approach relies fundamentally on price-sensitive behavior by consumers as a mechanism for controlling costs. The second analysis considers the effect of price on the choice between FFS Medicare and coverage through a managed care plan. This analysis addresses the specific question of how exposing beneficiaries to rising FFS costs will influence managed care enrollment.

Given the overlap in the data used, it is not surprising that the two analyses generate similar qualitative results. Both indicate that price is a significant factor influencing the plan choices of Medicare-eligible retirees, but that the effect is small. Comparisons with previous studies suggest that retirees are less sensitive to price than non-elderly individuals making health insurance decisions under similar circumstances.

The more willing consumers are to switch plans in response to an increase in price, the greater incentive there will be for plans to compete on the basis of price. Since it is not clear exactly how price sensitive consumers need to be in order to generate vigorous price competition, it is difficult to draw inferences from the results of this study about how vigorously health plans would compete in a managed competition–style Medicare program. Nonetheless, the pronounced differences between Medicare-eligible retirees and active employees suggest that caution should be taken in extrapolating from the experiences of employer-sponsored programs for the non-elderly employed to the Medicare program.

A potential explanation for the lower price sensitivity of retirees is that, for reasons related to age, they face higher switching costs than the employed non-elderly. However, regressions that allow the effect of price to vary with age among active employees and retirees provide only limited evidence that price sensitivity declines with age. Switching regression results indicate that active employees approaching retirement age (55 and older) behave more like younger active employees than younger retirees (72 and under). Similarly, FFS demand regressions suggests that the effect of price declines with age, albeit gradually.

The differences between active employees and retirees, therefore, may be related to the Medicare program itself or to retirement. For example, the cost of obtaining information on competing plans may be higher for retirees. Whereas employees may base their assessment of the quality of other plans on the experiences of coworkers in those plans, retirees will have less day-to-day contact with each other or with active employees, making such word-of-mouth information less prevalent. Also, the administrative tasks associated with switching plans—such as paperwork, phone calls, and so forth—may be more burdensome to the retired elderly.

In addition to these sources of switching costs, which may be relevant to retirees generally, the result of this study may be related to unique features of the geographic markets in which UC retirees are located. UC campuses are in areas that have among the highest rates of Medicare HMO penetration in the country, and even before the price of FFS coverage increased, a high fraction of UC retirees were in HMOs. Wisner, Feldman, and Dowd (1994) note the belief within the HMO industry in another mature managed care market, the Twin Cities, that the market for Medicare HMOs may become saturated at high levels of penetration, causing prospects for further penetration to be limited. According to this argument, the UC retirees who had resisted joining an HMO up until 1993 were those with the strongest preference for FFS

coverage, and the UC retirees with the most elastic demand were already in HMOs. Thus, while the behavior of UC retirees may be representative of a steady state situation under managed competition, it may understate the short-run price responsiveness in markets where Medicare HMO enrollment is currently low. The greater effect of price at the UC's one New Mexico location, where HMO enrollment was initially low, provides some evidence of this. This suggests the need for more analysis using data on retirees in other markets.

Whether maintaining FFS coverage as an affordable option is a desirable objective from an economic perspective, the availability and cost of FFS coverage is likely to be a key issue influencing the political viability of competition-based Medicare reforms. While the experience of large employers suggests that managed competition can be effective in controlling costs, it also suggests that without effective means of risk adjustment, biased selection can cause instability and drive certain plans from the market (Price and Mays 1985; Luft, Trauner, and Maerki 1985; Shewry et al. 1996; Cutler and Reber 1998). Plans placing fewer restrictions on enrollees' choice of providers are particularly vulnerable to adverse selection. Again, the UC is a good case in point. Since the UC adopted a fixed-dollar contribution, the Prudential FFS plan for active employees has undergone a classic adverse selection death spiral. Between 1993 and 1997, its gross premiums increased by 227 percent and its enrollment fell by 93 percent.

During the period studied here, Prudential's Medigap plan did not suffer a similar fate. While employee contributions for the plan increased over time, this trend was due to decreases in the UC's premium contribution rather than to large increases in Prudential's gross premium. Indeed, Prudential's Medigap premiums fell between 1993 and 1996. The difference in mean age between Prudential Medigap enrollees and other UC retirees (the only measure of risk in the UC administrative data) increased only slightly between 1993 and 1996, from 3.4 years (76.1 versus 72.7) to 3.8 years (76.9 versus 73.1). This suggests that there is an important positive aspect of a low elasticity of demand by Medicare beneficiaries: less price-induced switching will likely result in greater program stability and less of a problem of biased selection. Because of the lack of data on the health status or medical care utilization, this study can say little more about the potential for biased risk selection in a choice-based Medicare plan, or about the best way to address such problems should they arise. The policy debate would benefit from further empirical research focusing on how the effects of price and health status interact to influence the health plan decisions of the elderly.

APPENDIX

Table A1: Summary Statistics for Plan Switching and FFS Demand Analyses

Dependent Variables	*Plan Switching Regressions (n = 16,104) Mean (std. deviation)*	*FFS Demand Regressions (n = 68,888) Mean (std. deviation)*
Switched plans, 1993–1994 (S)	.045 (.208)	
Chose FFS coverage (F)		.514 (.500)
Independent Variables		
Change in premium, 1993–1994 (ΔP)	33.91 (33.63)	
Premium for FFS plan (P^F)		56.80 (52.85)
Age	73.53 (7.14)	74.23 (7.04)
Male with single coverage	.120 (.325)	.126 (.331)
Female with single coverage	.375 (.485)	.387 (.487)
Non-Medicare family member	.145 (.352)	.127 (.333)
HMO in 1993	.430 (.495)	
S.F. Bay Area	.327 (.469)	.326 (.469)
UC Davis	.068 (.252)	.102 (.303)
UC Irvine	.055 (.277)	.056 (.230)
UC Riverside	.037 (.189)	.037 (.190)
UC San Diego	.074 (.262)	.075 (.264)
UC Santa Cruz	.017 (.129)	.018 (.132)
UC Santa Barbara	.034 (.181)	.041 (.198)

Continued

Table A1: Continued

Dependent Variables	*Plan Switching Regressions (n = 16,104) Mean (std. deviation)*	*FFS Demand Regressions (n = 68,888) Mean (std. deviation)*
Lawrence Livermore National Lab.	.119 (.324)	.123 (.329)
Los Alamos National Lab.	.088 (.283)	.043 (.203)

Note: Retirees from Los Alamos represent a larger share of the switching sample because they contribute only two years of data (rather than four) to the FFS sample. A Blue Shield FFS plan was offered at UC Davis in 1993, but was canceled in 1994. Enrollees in this plan are included in the FFS sample but not the switching sample.

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NOTES

1. Klemperer (1995) identifies 6 sources of switching costs. Most apply to decisions relating to health insurance and are discussed in some detail in an earlier paper by Neipp and Zeckhauser (1985). Samuelson and Zeckhauser (1988) point to health plan choice as a decision that is subject to "status quo bias."
2. There is a fairly large literature on the demand for Medicare supplemental insurance. See Vistnes and Banthin (1997/1998) for a full set of citations.
3. Roughly four out of five Medicare recipients carry supplementary private insurance. An additional 7 to 8 percent qualify for supplemental Medicaid coverage by virtue of their low income (Morrisey 1993).
4. See Dowd, Christianson, Feldman, et al. (1992) for a good discussion of this issue.
5. All have a $5 copayment for physician office visits and no copayment for hospitalization. Copayments for prescription drugs range between $5 and $7.
6. See Morrisey (1993) for a description of the different ways the benefits of employer-sponsored plans are coordinated with Medicare.
7. The UC menu also includes UC Care, which was a PPO until 1995 and a "three-tiered" POS plan thereafter. For Medicare beneficiaries, UC Care offers "carve-out" coverage which is significantly less generous than that of Prudential High Option. As a result, when Prudential was free to Medicare beneficiaries, fewer than one percent was in UC Care.
8. The monthly out-of-pocket premium (OOP) paid by a UC retiree is given by the formula: OOP = max [0, gross plan premium + Part B premium − UC contribution].

9. In 1993, two plans required premiums contributions for Medicare coverage and one required a contribution for Medicare-eligible retirees with non-Medicare dependents. As a result, 8 percent of the retirees in my sample paid out-of-pocket premiums (i.e., paid a portion of their Part B premium) in that year.
10. Since nearly 90 percent of retirees have either single or two-party Medicare coverage, Table 1 presents premiums for those coverage categories only. Premiums tend to be higher for couples in which one spouse is not Medicare-eligible (13 percent) and families with dependent children (1 percent). In my analysis I drop retirees with dependent children, though the results do not change materially when they are included.
11. An additional econometric issue related to ΔP is that the variable takes the same value for all individuals in a particular plan or coverage category. As Moulton (1986) has shown for models with continuous dependent variables, failing to take account of the heteroscedasticity induced by this "clustering" will result in estimated standard errors that are biased downward. In estimating Equation 1, I use a conventional probit specification to estimate the model and adjust the standard errors accounting for the clustering using the method proposed by White (1980).
12. Price sensitivity likely varies across individuals in different states of health. If health status data were available, it would be interesting to investigate interactions between health status and price. The lack of such data, however, does not imply that the estimated effect of price is biased, but that it represents an average effect over all health states.
13. Long, Settle and Wrightson (1988) control for price and prior medical care utilization, and find no significant effect of the latter on health plan disenrollment. Riley, Feuer and Lubitz (1996) compare the HMO disenrollment of beneficiaries with and without cancer. They find that beneficiaries diagnosed with cancer prior to joining an HMO were more likely to disenroll than those without cancer, but found the opposite effect for those diagnosed after joining an HMO. Overall, these differences between types of cancer patients are offsetting, leaving no difference between beneficiaries with and without cancer.
14. In unreported regressions, I dropped couples in which one spouse was not Medicare-eligible. The results for these regressions are very similar to those in Table 2.
15. As would be expected, retirees living out-of-state are disproportionately enrolled in the FFS plans. This exclusion reduces the full sample by roughly 10 percent (1,759 observations) and reduces the HMO-only sample by less than 2 percent (121 observations). Dropping these observations has a minimal effect on the regression parameter estimates. The estimated price coefficients are slightly larger in the in-state sample than in the full sample, though not significantly so.
16. See Buchmueller and Feldstein (1997) and studies cited therein.
17. While more controls are available for active employees, for the sake of comparability, I use the same specification for actives as for retirees. Results for models with the fuller set of controls are reported in Buchmueller and Feldstein (1997).
18. Another possible difference between the active and retiree samples is that the

former includes observations choosing family coverage, whereas the latter does not. To test for possible differences in price sensitivity across coverage tier I interacted ΔP with a variable indicating single coverage. For both samples, the coefficient on this interaction was positive—indicating that single individuals are *more* price sensitive than couples—but not statistically significant.

19. As a test of the importance of individual fixed effects, I estimated linear probability models with and without a vector of dummy variables for each individual. Including or excluding these dummies had no material impact on the price coefficient.
20. This reduces the number of retirees from California by 8.3 percent and the New Mexico sample by 22.4 percent. Dropping out-of-state retirees has very little effect on the results for California, and increases slightly the estimated price elasticity for New Mexico.

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